

# **Childcare and Maternal Labor Supply – a Cross-Country Analysis of Quasi-Experimental Estimates from 7 Countries**

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## **Abstract**

The effect of childcare availability on maternal labor supply varies greatly by institutional context. The evaluation of interdependencies is difficult based on available single country estimates, as they differ both in their context and empirical methodology. We provide comparable quasi-experimental estimates of the childcare effect for 7 countries using harmonized data and the same method, based on country-specific kindergarten eligibility cutoffs and representative labor force data. The effect is shown to be dependent on the labor supply of mothers relative to those with older children, and the presence of constraints (leaves, job protection, norms, and job inflexibility).

**Keywords:** subsidized childcare, maternal labor supply, institutional context

**JEL Codes:** H24, J13, J22

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## 1. Introduction

Previous evidence on the effect of subsidized childcare availability on maternal labor force participation (LFP) suggests that the effect varies greatly among countries due to differences in their institutional contexts, including family policies, labor market flexibility cultural norms for instance (Cascio, Haider, and Nielsen 2015a; Vuri 2016). Yet there is little evidence on exactly how the childcare effect varies with these factors, and policymaking is mostly limited to general targets for childcare coverage – for example, the EU’s Barcelona Targets<sup>1</sup> – that are not linked to reforms of other potential limiting factors. This paper provides quasi-experimental estimates of the childcare effect for 7 EU countries (Austria, Czech Republic, France, Greece, Hungary, Italy, Slovakia) with varying institutional contexts, based on harmonized data and the same quasi-experimental methodology. The exogenous variation in childcare availability used for identification comes from country-specific kindergarten eligibility cutoffs around age 3 of children. We evaluate the country-level estimates in light of the cross-country variation in their family policies, labor market flexibility, and cultural norms, paying special attention to the point of estimation within each country. The results allow us to better evaluate differences in childcare effect estimates due to the institutional context by keeping the data quality, estimation method, and child age at measurement fixed across countries with varying contexts.

Methodologically, recent empirical research on the childcare effect has increasingly turned towards quasi-experimental methods based on policy changes or birthdate-based eligibility cutoffs. This is because childcare effect estimates may be biased due to omitted variables such as the economic development of regions, which affects the number of available childcare seats (through more abundant municipal resources) as well as the labor supply of mothers (through higher expected employment probabilities). While quasi-experimental estimates allow for better identification of the childcare effect due to the exogenous source of variation, it is important to note that they are local in nature, and therefore highly dependent on the estimation context, the age of the child at measurement, and the method of estimation.

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<sup>1</sup> The European Union set specific targets for its countries in 2002 and renewed them in the Europe 2020 Strategy, prescribing a 33% coverage rate for children under 3, and a 90% coverage rate for those between 3 and the mandatory school age by 2010 (EC 2013, 2008). While most previous estimates pertain to western countries with relatively supportive environments and already high maternal labor supply rates, little evidence is available from settings with very different institutional contexts, such as the Southern and Central-Eastern European countries. Since most of these countries are significantly behind in fulfilling the targets and expansion places a high financial burden on them, it is important to assess the expected labor market impact accurately given their context.

The comparison of single-country estimates is therefore not very informative regarding the causes of the cross-country differences of the childcare effect, as these may stem from methodological as well as contextual differences. A recent study highlights the relevant factors that most probably drive the differences (Cascio, Haider, and Nielsen 2015b), based on a review of single country estimates. First, the labor supply characteristics of mothers by child age are important. Specifically, the relative labor supply of mothers at a given child age compared to that of mothers with older children determines the magnitude of the childcare effect. The scope for policy to increase labor supply is higher when the targeted mothers' relative rate is low, since the difference reflects a qualified workforce that is potentially ready to enter the labor market in the short run. Second, interdependencies with other institutional elements - such as child-related leaves, labor market flexibility, and cultural norms - are also important, as these affect labor supply incentives and preferences. The effect of childcare expansion may be limited by financial disincentives, the lack of job protection and flexible work opportunities, or unfavorable views of maternal employment and institutionalized childcare. These factors impact the overall pattern of the maternal labor market reactivation process after giving birth, as well as the decisions of mothers at a given child age.

Quasi-experimental evidence from various countries seems to be in line with these points. No effect or a very small effect was found in the US (Cascio 2009; Fitzpatrick 2010), and France (Givord and Marbot 2015), where maternal employment rates of the treatment groups were already high. A more significant impact was found in Spain (Nollenberger and Rodríguez-Planas 2015), in 1996 Germany (Bauernschuster and Schlotter 2015), and in Hungary (Lovasz and Szabo-Morvai 2013) in settings where pre-treatment maternal employment rates were significantly lower. Some studies discuss the role of the leave system and cultural views in constraining the childcare effect (Givord and Marbot 2015; Nollenberger and Rodríguez-Planas 2015), and that of highly qualified mothers and the lack of childcare alternatives in magnifying it (Bauernschuster and Schlotter 2015). While quasi-experimental studies focus on estimating the causal effect of a single policy, a strand of policy literature analyzes the roles of various family policy elements and cultural norms in shaping maternal LFP based on cross-country comparisons (Boca, Pasqua, and Pronzato 2009; Cipollone, Patacchini, and Vallanti 2014). These show that the availability of childcare - especially under age 3 - the existence of job-protection and well-paid leave that are neither too short nor too long, flexible job opportunities, and cultural support for maternal employment are correlated with the relatively higher participation and working hours of mothers compared to childless women (Boeckmann, Misra, and Budig 2014). These findings provide a basis for

considering which institutional elements to include in the cross-country comparison of the childcare effect.

The uniform data and methodology used in our analysis allow a focused cross-country comparison of the role of the institutional context in determining the childcare effect. We estimate the causal effect of childcare availability for several countries utilizing the exogenous variation due to eligibility cutoffs for precise identification. The analysis combines representative harmonized European Labour Force Survey (EU LFS) data from 7 countries (covering 2005-2012), country-level information on birthdate-based kindergarten enrollment cutoffs and procedures provided by country experts and confirmed using further data sources, and country-specific institutional characteristics based on various data sources, such as the OECD Family Database, and the European Social Survey. As a first step, we discuss the countries' institutional contexts and document the country-level differences in the timing of mothers' labor market return after the birth of a child relative to major changes in family policy elements. The countries show distinct patterns that we use to group them into a few institutional categories, based on which we discuss our expectations of the childcare effect. Next, we estimate the effect on maternal LFP separately for each country, using an IV approach where date of birth serves as an instrument for childcare availability and potential seasonality biases are corrected for. We then compare these estimates in light of the institutional contexts, paying special attention to what is relevant to mothers' decisions at the exact point of estimation, or child age.

This study adds to the literature by being the first to provide harmonized and comparable quasi-experimental estimates of the effect of childcare on maternal labor supply for several countries that represent varying institutional contexts. Although the size and representativeness of the sample of countries analyzed is limited by the requirements of the estimation method, the existence of a cutoff, and data availability, the comparison reveals clear differences in the childcare effect between the groups of countries that point to specific policy implications. The EU countries included in the analysis vary enough in their institutional contexts to provide internationally policy relevant implications. The results show that the childcare effect is the highest in CEE countries, where at this child age, maternal participation is still relatively low compared to that of mothers with older children, and leaves with job protection are just ending. We find less evidence of an impact in Southern EU countries, where leaves end at a much earlier age, and maternal participation at older child ages is low due to mothers permanently exiting the labor market earlier on. Western EU countries show some impact, despite the relatively high maternal participation rates at this

child age. The results suggest that for childcare expansion to have a high impact on maternal labor supply in countries with relatively low childcare coverage and maternal participation, parallel reforms of the leave system, flexible work regulations, and shaping of public opinion are key.

## **2. The Role of the Institutional Context**

To outline how the institutional context affects the quasi-experimental estimates of this paper, we first give a brief overview of the previous empirical findings in the literature on the determinants of maternal labor supply. We review the available evidence on the role of each factor, and consider the possible interactions of childcare availability and other factors. We then discuss the differences in these measures among our sample of countries and those for which previous single-country quasi-experimental childcare effect estimates are available, and group them into more general categories. We derive hypotheses regarding their likely impact on the childcare effect, which we analyze empirically later on. Finally, we present figures depicting the timing of mothers' return to the labor market over the age of their youngest child, highlighting country-level differences, the relationship of maternal participation and the timing of relevant policy changes, and the points at which our estimations are carried out.

### **2.1. Key institutional characteristics**

Previous evidence on the effects of various family policies on maternal labor supply come from three main sources: quasi-experimental evidence, structural estimates, and cross-country analyses. Evidence on the effect of childcare availability from single countries is highly variable. One strand of studies focuses on structural models and generally utilizes regional and time variation for identification. Some support the existence of a childcare effect (Connelly 1992; Del Boca 2002; Haan and Wrohlich 2011; Kimmel 1992; Lokshin 2004), while others find little or no significant impact (Chevalier and Viitanen 2005; Chone, Leblanc, and Robert-Bobee 2003; Ribar 1995). Several recent studies use exogenous variation in childcare availability related to policy changes, or utilize eligibility cutoffs to identify the childcare effect. Some find a significant positive impact (Baker, Gruber, and Milligan 2008; Bauernschuster and Schlotter 2015; Berlinski and Galiani 2007; Bettendorf, Jongen, and Muller 2015; Gelbach 2002; Givord and Marbot 2015; Haeck, Lefebvre, and Merrigan 2015; Hardoy and Schøne 2015; Lefebvre and Merrigan 2008; Nollenberger and Rodríguez-Planas

2015), while others find no effect (Cascio 2009; Fitzpatrick 2010; Havnes and Mogstad 2011; Lundin, Mörk, and Öckert 2008). Cross-country comparisons also suggest that subsidized childcare availability under age 3 of children is strongly correlated with maternal labor supply (Boeckmann, Misra, and Budig 2014; Budig, Misra, and Boeckmann 2012).

Regarding the leave system, evidence suggests that both the lengths and the benefit amounts of the leaves available to mothers are important factors in determining maternal labor supply. Previous studies suggest that moderately long, well-paid leaves increase maternal LFP (Boeckmann, Misra, and Budig 2014; Keck and Saraceno 2013; Olivetti and Petrongolo 2017). Very short - or non-existent - leaves constrain the opportunities of women to reenter their jobs, and discourage women from higher income households to return to work. On the other hand, very long, low-paying leaves may lead - especially low-skilled - mothers to become detached from the labor market and the depreciation of their skills, as well as increased statistical discrimination against mothers and women (Boeckmann, Misra, and Budig 2014). In our cross-country analysis, we focus on a few aspects of leave policies: the length of paid leave (and job protection) available to mothers, and the amount of the benefit that is available to mothers during the leave. Previous available evidence on the interactions of leaves and childcare availability suggest that non-optimal leaves may restrict the effect of childcare expansion (Geyer, Haan, and Wrohlich 2015), and that the lack of childcare may limit the positive effect of leaves (Cukrowska-Torzewska 2015).

The flexibility of labor markets is also an important factor in determining mothers' labor market status, though it lies outside the direct realm of family policies. Empirical evidence so far mainly focuses on the effect of part-time work opportunities on maternal LFP. The employment rate of mothers with young children has been shown to be strongly correlated with the availability of part-time work opportunities: part-time work may provide mothers with a means to strengthen their attachment to the labor market and keep their skills up to date, while allowing for a more gradual separation from their child. Furthermore, the quality (related job protection, social benefits and earnings) of the available part-time jobs also has a significant impact (Del Boca 2002).

Cultural norms have also been shown to be strongly correlated with maternal outcomes, and unfavorable attitudes towards maternal labor force participation may limit the effectiveness of family policies. Studies seek to identify the effect of culture on maternal labor market outcomes in several ways. One study compares migrants with different cultural values, who live in the same economic and institutional setting, finding a significant impact (Fernandez 2007). Other studies use various available indices describing views on child

development and female employment, to show that they affect maternal outcomes (Budig, Misra, and Boeckmann 2012; Fortin 2005). The interdependencies of policies and norms have been discussed extensively in social policy studies (Pfau-Effinger 1998), however, the relationship is very difficult to identify empirically, and remains mostly unclear (Kremer 2007). Overall, evidence suggests that norms may limit the effectiveness of family policies (Budig, Misra, and Boeckmann 2012). A 2010 report of the European Commission (Mills et al. 2014) on the evaluation of the fulfillment and effectiveness of the Barcelona childcare targets also notes the importance of norms related to parenthood, institutionalized childcare, and parental preferences at the country level, and the need for these norms to be shaped through raising public awareness.

Finally, the role of alternative childcare options, including private and informal care, is also important to consider. Private childcare plays an important role in some western European countries, but is very scarce and unaffordable to most people in the CEE countries. On the other hand, informal childcare is highly common in several of the countries we study, particularly the CEE and southern European countries, due to the presence of a large inactive elderly population and the relatively low mobility in these countries. Informal childcare may be important in allowing mothers of younger children to work, especially when formal childcare is rationed (Ghysels 2011; Posadas and Vidal-Fernández 2012). This may be especially true in countries where views are generally unsupportive of institutionalized care at young child ages, such as the CEE countries (Saxonberg and Sirovátka 2006).<sup>2</sup>

Though each of these factors have been shown to play a role in determining maternal participation, there is little direct evidence on their interactions with each other, and with childcare availability in particular. Budig and coauthors (2012) show that cultural attitudes moderate the impact of policies on women's earnings across countries. Cukrowska-Torzewska (2015) estimates the effect of various policy measures on maternal employment and wages, based on individual level data from 28 European countries, allowing for the interaction of childcare availability and leave policies. The findings indicate that the impact of leave is dependent on childcare availability: long maternity leaves combined with high childcare coverage lead to a higher gap in the employment of mothers and non-mothers compared to settings where the coverage is low. The study of Geyer and coauthors (2015)

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<sup>2</sup> The former socialist countries were characterized by a very well developed childcare system, with relatively high nursery school coverage under age 3, which was dismantled following the transition. Nursery schools, however, were not considered pedagogical institutions, but rather healthcare ones, and were not regarded positively by the population.

from Germany analyzes the combined effect of the expansion of subsidized childcare and a simultaneous reform of the leave system that increased the benefit amount but reduced the length available. It does so using a structural model, as the exogenous variation in the two factors did not occur at the same time. The study finds that a combination of parental leave benefits and subsidized childcare can increase maternal labor supply significantly. In this study, we focus on the estimation of the effect of childcare availability, and evaluate how the other relevant factors may impact this effect based on the cross-country comparison of a set of high internal validity, but very local childcare effect estimates.

## **2.2. Estimation context**

### **2.2.1. Country-level characteristics**

The sample of countries included in our analysis is determined by data availability and the existence of kindergarten eligibility cutoffs that are necessary for the identification strategy. The final set of 7 EU countries differ significantly in key aspects of their institutional environments, which are likely to influence the effect of childcare availability – specifically, kindergarten eligibility - on maternal labor supply. Table 1 summarizes the factors described in the previous subsection that play a role for maternal LFP. The countries in our analysis are included in the table, as well as further countries from which quasi-experimental evidence is available. The countries are grouped into categories by geographical regions, which are characterized by certain sets of institutional traits that are likely to impact maternal LFP similarly. At the same time, there is variation in the key factors among countries within these regions, which we also discuss.

The CEE countries in our sample exhibit some strong similarities due to their shared socialist institutional and historical heritage (Lovász 2016). CEE countries generally have very low maternal participation rates below age 3 of children, but relatively high rates at older child ages. Formal childcare enrollment shows a similar pattern, with the lowest rates at age 2 of children among the EU countries. CEE countries provide very long leaves to mothers (parents), with job protection and high amounts of cash benefits even at age 3 of children. Family policies therefore clearly encourage mothers to stay home until around this age. The low availability of part-time jobs is also not conducive to mothers' earlier return to work, and informal childcare plays a relatively important role due to the presence of a large inactive elderly population. Views are less supportive of maternal employment compared to western European countries, despite the socialist rhetoric of gender equality, or as a response to it.



Maternal participation rates in the Southern European countries are higher under age 3 of children compared to CEE countries, but their increase is relatively minor as children grow older. Childcare enrollment rates are higher at age 2 as well, and, in the case of Spain, relatively high overall within the EU. Leaves for mothers are much shorter, and very short - 16 weeks - in the case of Spain, and cash benefits received at age 3 of children are significantly lower than in CEE. Part-time work makes up a higher proportion of jobs compared to the CEE, but still lower than what is seen in western EU countries. The southern EU countries are generally characterized by traditional cultural views and gender norms. Although their family policies do not explicitly encourage mothers to stay home, the short leaves, coupled, in some cases, with low childcare availability, and the unsupportive norms eventuate that many mothers do not return to work after having a child, and fall out of the labor market completely.

The countries in the Western EU group are diverse in many aspects. Germany and Austria are generally traditional in cultural norms and were historically less supportive of female employment. However, they made significant changes aimed at increasing maternal employment, including the expansion of childcare under age 3, and are characterized by relatively high maternal employment and a high availability of part-time jobs. France and Sweden represent some of the western countries that are most known for supporting gender equality, with Sweden often being cited as a role model in terms of policies supporting maternal employment and gender equality. These western EU countries exhibit the highest maternal participation rates and childcare enrollment rates below age 3 of children, and which are further linked to very flexible labor market opportunities.

The US and Canada are included in the table due to the significant strand of empirical evidence (see Col. 12 in Table 1) on the childcare effect available for these countries. They are generally characterized by relatively high maternal employment under age 3 of children. State support available to mothers is significantly lower, with low formal childcare enrollment at both age 2 and 3, low cash benefit amounts, and very short (or non-existent) leaves. On the other hand, these countries are generally characterized by liberal norms, supportive of gender equality and female employment. The final columns of the table summarize the available quasi-experimental estimates seen in previous studies and the countries analyzed in this paper. The table shows that most of the empirical evidence comes from Western European countries or North America, with much less evidence from Southern EU or CEE countries, which have very different institutional contexts, and therefore, likely different potential effectiveness of childcare expansion.

Table 1: Institutional characteristics of the countries in the estimation sample and previous studies

Region	Country	Maternal employment rate (%) at child age ...			Childcare enrollment (%) at child age ..		Informal childcare	Child-related leaves			Labor market flexibility	Preferences / norms	Literature				
		0-2	3-5	6-14	2	3		Leave for mothers (weeks)	Total leave - average replacement rate (%)	Cash benefits at age 3 (%)			Share of female part-time in employment (%)	Child suffers	Reference	Age of child at the point of estimation	Effect
		(1)	(2)	(3)	(4)	(5)		(6)	(7)	(8)			(9)	(10)	(11)	(12)	(13)
CEE	Czech Republic	20	70	87	7.0	41.9	24.3	110	51.1	16.4	10.0	.	.	.	.		
	Hungary	12	63	75	16.7	60.2	18.9	160	44.5	23.0	9.0	54.7	Lovasz-Szabo-Morvai, 2013	3	+		
	Slovakia	15	56	80	6.7	46.4	16.3	164	32.0	17.7	6.0	44.4	.	.	.		
Southern EU	Spain	55	57	59	60.4	84.4	9.0	16	100.0	4.0	25.0	46.5	Nollenberger and Rodriguez-Planas, 2015	3	+		
	Greece	50	54	59	28.7	49.1	32.5	43	53.9	5.2	12.5	65.3	.	.	.		
	Italy	51	53	56	38.4	81.5	18.0	48	52.7	5.2	31.6	61.8	.	.	.		
Western EU	Austria	67	74	82	26.5	54.7	18.7	60	85.3	12.6	46.0	54.8	.	.	.		
	France	61	74	79	58.0	86.2	7.9	42	44.7	12.1	.	41.0	Givord and Marbot, 2015	pre-school	0		
	Germany	52	70	78	.	.	6.1	58	73.4	15.1	47.1	49.8	Bauernschuster and Schlotter, 2015	3	+		
	Netherlands	75	75	78	.	.	20.2	16	100	5.8	75.6	44.4	Bettendorf et al., 2015	0-12	0		
	Sweden	.	.	.	84.5	87.6	.	60	63.4	7.8	36.3	31.0	Lundin et al., 2008	1-9	0		
Americas	Canada	67	72	79	.	46.0	.	17	52.8	.	26.0	57.3 <sup>(15)</sup>	Baker et al., 2008	0-4	+		
		67	72	79	.	46.0	.	17	52.8	.	26.0	57.3 <sup>(15)</sup>	Haeck et al., 2015	1-4	+		
		67	72	79	.	46.0	.	17	52.8	.	26.0	57.3 <sup>(15)</sup>	Haeck et al.,	5	0		

													2015		
		67	72	79	.	46.0	.	17	52.8	.	26.0	57.3 <sup>(15)</sup>	Lefebvre and Merrigan, 2008	4	+
	United	56	62	70	.	66.0	.	0	0.0	4.3	17.0	.	Cascio, 2009	5	0
	States	56	62	70	.	66.0	.	0	0.0	4.3	17.0	.	Fitzpatrick, 2010	4	0

- (1) Employment rate of mothers with youngest child aged 0-2, %. Source: OECD Family database, <http://www.oecd.org/els/family/database.htm> LMF1.2.C. Maternal employment rates by age of youngest child (2013)
- (2) Employment rate of mothers with youngest child aged 3-5, %. Source: OECD Family database, <http://www.oecd.org/els/family/database.htm> LMF1.2.C. Maternal employment rates by age of youngest child (2013)
- (3) Employment rate of mothers with youngest child aged 6-14, %. Source: OECD Family database, <http://www.oecd.org/els/family/database.htm> LMF1.2.C. Maternal employment rates by age of youngest child (2013)
- (4) Own calculations using EU-SILC data for years 2005-2012 based on the methodology of OECD Family Database.
- (5) Own calculations using EU-SILC data for years 2005-2012 based on the methodology of OECD Family Database.
- (6) Own calculations using EU-SILC data for years 2005-2012 based on the methodology of OECD Family Database.
- (7) Full-rate equivalent total paid leave for mothers (weeks). OECD Family database, <http://www.oecd.org/els/family/database.htm> PF2.1.A. Summary of paid leave entitlements (2015)
- (8) Average replacement rate (%): proportion of previous earnings replaced by the benefit over the length of the paid leave entitlement for a person earning 100% of average national earnings. OECD Family database, <http://www.oecd.org/els/family/database.htm> PF2.1.A. Summary of paid leave entitlements (2015)
- (9) Cash benefits and tax breaks at the child age of 3, relative to the median working age income, %. Source: Source: OECD Family database, <http://www.oecd.org/els/family/database.htm>
- (10) Part-time employment as a % of all employment, 20-64 year-old females, 2013. Data source: Eurostat, <http://ec.europa.eu/eurostat/data/database>
- (11) European Values Study, <http://www.europeanvaluesstudy.eu/>; Pre-school child suffers with a working mother. 0: Strongly disagree; 100: Agree strongly. Sample: 20-50 year-old females, waves 1999-2001 and 2008-2010
- (14) "0": No significant effect or very small effect; "-": Significant negative effect; "+": Significant positive effect
- (15) A pre-school child is likely to suffer if both parents are employed (0 - disagree strongly; 100 - agree strongly, rescaled) (1999). Source: Canadian Attitudes on the Family, <http://www.imfcanada.org/sites/default/files/Canadian%20Attitudes%20on%20the%20Family.pdf>

### **2.2.2. Timing and the point of estimation**

Our study contributes to the discussion regarding childcare policies by providing comparable estimates from countries with a wider variety of settings. The comparison of the estimates also needs to take the point of estimation into account, as the incentives and constraints mothers face, and thereby the magnitude of the childcare effect, differs not only by country, but also by child age within countries. Most of the studies from Western Europe and the US found little or no evidence of a childcare effect, measuring at child ages (Table 1, Col. 13) at which maternal participation is already high relative to that of mothers with older children or females. In such settings, the potential for childcare policies to have an impact is low due to the already high rates. The three previous studies from settings where maternal participation is relatively low at the point of estimation, from Spain (Nollenberger and Rodríguez-Planas 2015), Hungary (Lovasz and Szabo-Morvai 2013), and Germany (Bauernschuster and Schlotter 2015), however, all point to a significantly higher childcare effect.

Figure 1 depicts the country-level variation in the timing of mothers' return to the labor market following the birth of their child for the sample of EU countries analyzed in this study, based on the EU-LFS data used in the analysis. It shows that the dynamics of mothers' return to the labor market as a function of the age of their youngest child is rather dispersed. The CEE countries (Czech Republic, Hungary, Slovakia) show the lowest rates under age 3 of children – in line with institutions that do not support employment under age 3 – but high rates at older child ages. The evolution of maternal LFP appears to be closely correlated with the evolution of childcare enrollment, and negatively correlated with the amount of cash benefits received related to the child. Maternal participation rates in the southern countries (Greece, Italy), on the other hand, are relatively stable as children age, with no significant increase when childcare enrollment increases. The two western countries in our sample (Austria, France) show higher maternal employment rates at all child ages, with a small increase around the time when childcare enrollment increases.

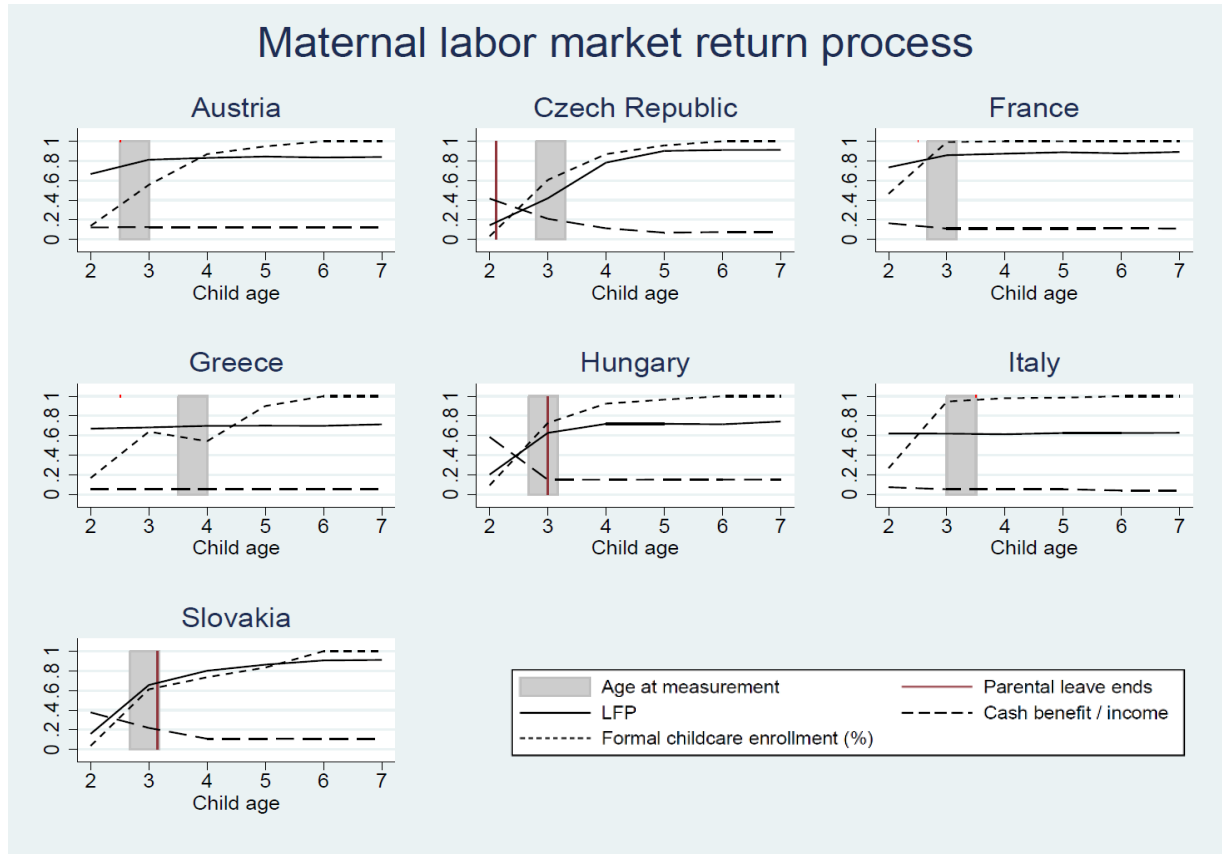
The figure also highlights the timing of important changes in the most relevant institutional factors, as well as the point of estimation of the childcare effect for each country in our analysis. Based on the country and child age level institutional characteristics and the point of estimation, we can form some hypotheses regarding its expected magnitude. The magnitude of the childcare effect is dependent on the characteristics at the point of measurement, but also on traits relevant at earlier and later ages, i.e. the overall characteristics

of the institutions. The childcare effect in a given country at a given child age is likely to be higher if (a) there is an underutilized, qualified, and willing workforce of mothers available, i.e. if maternal participation is still low at the point of estimation relative to the long-run “potential” rate (LFP at higher child ages), (b) mothers are able to return to protected jobs, and are not financially or culturally dis-incentivized from doing so, and (c) mothers are able to return gradually with the aid of part-time jobs.

The location of the cutoff point and thus the point of estimation in terms of child age is relatively stable across countries. Our analysis focuses on kindergarten eligibility cutoffs around age 3 of children, which exist in several EU countries as a border between the lower coverage nursery schools and the high coverage kindergartens (Table 1, Cols 4-5). However, the labor market characteristics of mothers at the point of estimation – including their relative labor supply compared to their long-run rate - depend on the institutional system of each country, affecting the interpretation and external relevance of our estimates. For the CEE countries (Czech Republic, Hungary, Slovak Republic), leave is just ending or ended recently and cash benefits drop significantly at the child age when the eligibility cutoff affects mothers. At this point, maternal participation rates are still well below those of mothers with older children, but as children age, they rise markedly. This suggests that in our estimations, increased childcare availability around age 3 is likely to have a high impact in these countries, as there is a readily available maternal workforce, financial incentives and cultural norms encourage mothers to return to work around this age, and their jobs are still protected, making their return easier.

In the Southern EU countries (Italy, Greece), leave and job protection have ended long before the point of estimation, and maternal participation is at a low level, though not relative to rates seen for mothers of older children. Childcare availability is therefore expected to have a lower impact, since mothers who were willing and able to return to work had likely already done so using informal childcare arrangements and flexible work opportunities, so the willing and able unutilized workforce is likely to be smaller. For the Western EU countries in our sample, leaves of medium length have already ended as well, and cash benefits are also low around age 3. Although maternal participation rates are already high, they still show some growth after age 3 of children, suggesting that some unutilized workforce is still present at the child age when they are included in our analysis. Childcare availability may therefore still have an impact despite already high participation rates, though the magnitude is expected to be lower compared to CEE countries with larger potential workforces.

Figure 1: Maternal return to the labor market following childbirth by country (2005-2012)



Note: The LFP rates are calculated from EU-LFS data, and the child age reflects the age of the youngest child in the household. Formal childcare enrollment rates are based on EU-SILC information and the calculation method follows that of the OECD Family Database. The information regarding cash benefits comes from the OECD Family Database and reflect the total family cash benefit spending of each country at a given child age as a proportion of the median working-age household income. The graphs are based on yearly data in terms of child age.

### 3. Data and Methodology

The analysis of the effect of childcare availability on the participation of mothers is based on individual-level EU LFS data from 7 countries. The sample of countries is determined by (a) the availability of birth date and age information on the youngest child of mothers in the EU LFS dataset, and (b) the existence of a kindergarten eligibility cutoff. We first describe the details of these two conditions and the resulting estimation sample, then describe the instrumental variable approach used to estimate the childcare effect.

#### 3.1. Dataset and variables

To apply our empirical approach, we utilize individual-level information on mothers' labor market activity and family status. In the EU LFS dataset, the exact day and month of birth of the youngest child are excluded for data security reasons, only the age (in years) of

the youngest child in the household is observed, thus the quarter of birth is not directly observable in the data. However, when we observe at least 4 quarters of observations in a row, we can infer the quarter of birth by observing when their age changes. For the countries included in the analysis, we are able to construct a stochastic panel of at least 4 quarters by linking household observations over quarters. We utilize a linking procedure to link household observations over time, for each country where the data was originally collected as a panel dataset. Linking is based on exact matches (or logical increases/decreases) of 56 variables describing the household level characteristics, household composition, and individual characteristics of certain members of the household, like year of completing highest level of education. We then derive the birth month of the youngest child by observing in which quarter (wave) their age changes, and assigning the interview month when the older age is first observed as the quasi birthdate.<sup>3</sup> It is only possible to construct such panel data for countries where data was originally collected as a household panel and where the structure of the database is suitable,<sup>4</sup> which limits the possible number of countries included.

Once birth dates are derived for the youngest child observed in each household, we identify the mothers of the youngest child using the parent codes available in the dataset. We limit our sample to these mothers, those aged 20-50, and those who were born in the given country and are therefore more likely to share the country-specific beliefs and values. We utilize data for the years 2005-2012, for which the key variables are observed and harmonized for all of the countries in our sample. For each mother, we observe individual-level labor force participation, employment, other labor market characteristics, demographic characteristics such as age and education, family status and characteristics of their spouse, household level characteristics regarding their composition and dwelling, and the region of their household in some countries. We also observe the birthdate of their youngest child, which is used to classify them into treatment (kindergarten eligible) and control groups, as described next. Table 3 depicts some descriptive statistics of the resulting dataset for the overall sample and the treatment and control groups respectively.

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<sup>3</sup> Households differ in their month of observation within the quarter. When we observe a change in the youngest child's age between two quarters, we know that the birthdate of the child lies between the two interview months. We assign the month of the latter interview to the child as the month of birth, so the month of birth of each child is either in the month assigned, or in the two previous months. As a result, birth dates are known to a quarterly precision, and we have birth data with a monthly frequency. We take this into consideration when determining our treatment and control groups around the eligibility cutoff by excluding the 3-month birth date groups overlapping treatment and control birth periods.

<sup>4</sup> For instance, some countries submit samples of the national LFS quarterly, but some of them submit only once a year which prevents the linking process.

Table 2: Descriptive statistics of the sample by country (2005-2012)

	Austria		Czech Republic		France		Greece		Hungary		Italy		Slovakia	
	(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)
Lower secondary (%)	0.090	0.099	0.080	0.105	0.176	0.170	0.212	0.200	0.240	0.256	0.375	0.366	0.105	0.090
<i>T-test</i> <sup>(3)</sup>	(0.619)		(0.184)		(0.680)		(0.536)		(0.458)		(0.885)		(0.514)	
Upper secondary (%)	0.715	0.723	0.782	0.753	0.419	0.439	0.493	0.514	0.599	0.564	0.518	0.504	0.720	0.754
	(0.791)		(0.283)		(0.292)		(0.399)		(0.155)		(0.834)		(0.315)	
Tertiary (%)	0.194	0.178	0.137	0.142	0.405	0.391	0.295	0.286	0.162	0.181	0.107	0.130	0.175	0.156
	(0.496)		(0.842)		(0.456)		(0.706)		(0.311)		(0.590)		(0.504)	
Widowed (%)	0.056	0.042	0.077	0.091	0.033	0.050	0.031	0.028	0.090	0.078	0.000	0.057	0.040	0.042
	(0.305)		(0.412)		(0.025**)		(0.646)		(0.379)		(0.010**)		(0.921)	
Single (%)	0.283	0.277	0.144	0.158	0.407	0.390	0.008	0.010	0.205	0.204	0.045	0.057	0.093	0.118
	(0.845)		(0.568)		(0.368)		(0.686)		(0.941)		(0.671)		(0.285)	
Married (%)	0.662	0.681	0.779	0.751	0.560	0.560	0.961	0.963	0.705	0.719	0.955	0.886	0.868	0.841
	(0.511)		(0.309)		(0.994)		(0.837)		(0.546)		(0.052*)		(0.325)	
Mean age of mothers (yr)	33.192	33.090	31.598	31.909	33.091	33.749	34.599	34.419	31.814	31.618	33.250	34.358	30.413	31.360
	(0.756)		(0.313)		(0.001***)		(0.484)		(0.478)		(0.074*)		(0.018**)	
Age of youngest child <sup>(4)</sup>	2.869	2.903	2.985	3.000	2.869	2.897	3.410	3.392	2.930	2.952	2.750	2.750	2.891	2.919
	(0.190)		(0.005***)		(0.002***)		(0.400)		(0.000***)		(.)		(0.186)	
LFP (1Q)	0.715	0.753	0.282	0.377	0.773	0.799	0.674	0.694	0.415	0.465	0.598	0.569	0.328	0.388
	(0.165)		(0.001***)		(0.105)		(0.388)		(0.043**)		(0.653)		(0.104)	
N	1043		985		2744		1662		1605		235		689	

Notes: (1) Control group mean (2) Treatment group mean (3) P-value of the two-sample t-test (4) In the Italian sample there is no variation of the youngest child's age, thus the t-test cannot be performed.



### **3.2. Kindergarten eligibility cutoffs**

To determine which have a birthdate-based kindergarten eligibility cutoff, we surveyed experts from each potential country, asking for detailed information regarding kindergarten enrollment rules, practices, and their changes over time. Experts were compensated for their contribution to ensure a high quality of answers.

The issue of determining cutoffs for each country and each year in our sample is not straightforward for several reasons. First, in most countries, actual practice differs from what is required in the legislature, or the law only states minimum requirements, and what is realized depends highly on the supply and demand for childcare spots in the given location at the given time. For example, in Hungary, the legislature states that children born prior to September 1st in each year must be accepted into kindergarten, while those born after may be accepted if spots remain available. In a previous study (Lovasz and Szabo-Morvai 2013), more detailed enrollment data is used to show that the effective cutoff during the time period studied (1998-2009) was actually January 1st: children born up to that date were generally accepted into kindergarten, while those born after had to wait until next September. Experts were asked to discuss such flexibilities in the legal cutoff specifications as well as real-life practices, but the determination of the exact effective birthdate cutoff that is needed to provide exogenous variation in childcare availability remained difficult. Second, information on the current legislation and especially on current real-life practices is easier to obtain than retrospective information. Although experts were asked to provide enrollment details by year, such data is also likely to be less precise. Finally, effective cutoff dates are also likely to vary regionally. For example, in large cities, where demand for childcare is relatively high, spots are likely to fill up closer to the legal cutoff, while in low population areas, children born much later may be able to enroll. We cannot account for such variation in our definition of the cutoff.

To minimize the chance of cutoff date misspecification, we turn to external data sources to verify the information received from the country experts. There is no individual-level information on actual childcare usage in the EU LFS dataset, which could be used to directly determine the eligibility cutoff, which is also why we are only able to derive reduced form estimates in our analysis. We do, however, verify the eligibility cutoffs for as many countries as possible, using some further data sources with birthdate and childcare enrollment-relevant information. First, we use EU-SILC data on actual childcare enrollment and quarter

of birth information to compare the enrollment rate means by birthdate group over various ages of children. The categorization of birthdate groups based on birth quarter is not exactly the same as what we use in our analysis, and not all quarters are observed in every country, which limits the test. However, the comparison of the birth quarters available does provide evidence on the existence of any discontinuity in childcare enrollment. Second, we use EU-LFS data on the mother's response to a survey question asked only from those not participating in the labor market, regarding the reason for their inactivity. We analyze whether there is a significant difference by birthdate group (defined the same way as in our estimation) in the likelihood of "looking after children or incapacitated" being given as the main reason. This measure also serves as a proxy for childcare enrollment, however, if we observe significant differences in the response rates by birthdates – even when controlling for individual and household characteristics – it also provides some indirect evidence verifying the existence of a cutoff.

In both tests, we assess whether significant differences exist at the cutoff expected based on the country expert responses, as well as whether differences can be seen after any other calendar dates. Table 3 summarizes the information used to determine the eligibility cutoffs for each country, including the results of these tests. For the EU-SILC enrollment test, the table gives the largest value seen among the mean differences in enrollment rates between birth groups over the child ages of 2 to 5. For most countries, we do not see significant differences elsewhere indicating other cutoffs, except for Hungary, which also shows a smaller but still significant difference at January 1st. In case of the EU-LFS test, we run a regression of the treatment and the control group defined by the effective cutoff date on the reason for not participating on the labor market. In Table 3 we report the coefficient and significance from this regression.

The effective cutoff date refers to the cutoff used in our analysis, which is verified by at least two of the three independent data sources. The effective cutoff date differs from the legal minimum requirement in several cases, due to the possibility of enrolling further children if capacity allows. In the case of several countries, the tests show that children with birthdates after September 1st could enroll in kindergarten up to the later birthdates of January 1st (Czech Republic) or March 1st (Hungary, Slovakia, France).<sup>5</sup> It is important to emphasize

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<sup>5</sup> The March 1 cutoff corresponds to having turned 2.5 years old by September 1<sup>st</sup>, which, in the case of Hungary, has been an increasingly common rule of thumb used by kindergartens in admissions, leading to a change in the law in 2010 specifically allowing it.

that, although it is key to the quasi-experimental analysis, determining the country-specific childcare enrollment cutoffs is clearly not straightforward. In most countries, the exact legislative rules can be overridden for the system to remain flexible enough and to account for regional and timely variations in childcare demand and supply. We therefore ensure the reliability of the cutoff specification by using the three independent sources of information, expert information, EU-SILC and EU-LFS data and include only countries in which at least two of the three sources confirm the cutoff date. Despite the benefit of further variation in institutional context, some potential countries that meet the other data requirements are excluded from the analysis due to unverifiable or ambiguous cutoffs.

*Table 3: Country cutoff details and sources of information*

Country	Effective cutoff date	Enrollment date	Expert information: the child can be enrolled in kindergarten if she is ...	EU-SILC		EU LFS Reason for inactivity is childcare: Coefficient of T (P-value)
				Birth quarters where largest enrollment difference exists	Mean difference in enrollment rates (P-value)	
<b>Czech Republic</b>	January 1	September 1, prior to cutoff	3 years old by Sept 1 <sup>st</sup> , or younger if spots available	q4-q1	0.23 (0.00)	-0.09 (0.01)
<b>Hungary</b>	January 1 (2005-2009), March 1 (2010-2012)	September 1, prior to cutoff	3 years old by Sept 1 <sup>st</sup> , or younger if spots available, 2.5-year-old from 2010	q4-q1 q1-q2	0.12 (0.00)	-0.08 (0.02)
<b>Slovakia</b>	March 1	September 1, prior to cutoff	3 years old by Sept 1 <sup>st</sup> , or younger if spots available	N/A	N/A	-0.05 (0.06)
<b>Greece</b>	May 1	September 1, after cutoff	30 months old on Sept 1 <sup>st</sup> , or younger if spots available	q1-q3	0.21 (0.01)	-0.02 (0.19)
<b>Italy</b>	May 1 (2005-2006) September 1 (2007-2012)	September 1, after cutoff	to 2006: 3 years old by May 1 <sup>st</sup> , after 2006: 3 years old by September 1 <sup>st</sup>	q2-q4	0.24 (0.00)	-0.04 (.12)
<b>Austria</b>	May 1	September 1, prior to cutoff	30 months old on Sept 1 <sup>st</sup> , or younger if spots available	q1-q2	0.15 (0.01)	-0.05 (0.06)
<b>France</b>	March 1	September 1, prior to cutoff	3 years old by January 1 <sup>st</sup> , or younger if spots available	N/A	N/A	-0.01 (0.09)

### 3.3. Empirical specification

In our empirical analysis, we estimate the childcare effect for each country, based on an eligibility cutoff-based IV methodology similar to what was used previously in an analysis focusing on Hungary (Lovasz and Szabo-Morvai 2013). The basic idea of the methodology, inspired by Angrist and Krueger (1991), is to use the birthdate of the child for the identification of the childcare effect. Mothers whose children are born before the kindergarten eligibility cutoff are eligible for kindergarten, while those born after the cutoff are only eligible for nursery school, which has significantly lower coverage in each country included in the analysis (cf. Table 1 Cols 4 and 5). Birthdate is therefore highly correlated with childcare availability, and, if it can be considered random – which we will discuss further in the next section – it is exogenous to maternal labor supply. By using the birthdate as an instrument, we can remove bias due to endogeneity in childcare availability, which may arise due to omitted variables such as the economic development of regions, which affects the number of available childcare seats (through more abundant municipal resources) as well as the labor supply of mothers (through higher expected employment probabilities).

Due to small sample sizes and the above-mentioned constraints on birthdate data, we define the instrument used in our analysis as follows. The treatment variable is defined as:

$$T_{cyi} = \begin{cases} 1 & \text{if } cutoff\ date - 5\ months \leq b_i \leq cutoff\ date \\ 0 & \text{if } cutoff\ date \leq b_i \leq cutoff\ date + 5\ months \end{cases} \quad (1)$$

where  $b_i$  is the youngest child's date of the birth, and the cutoff date varies by country (see Table 3). Because of the huge differences between availability of kindergarten and nursery, treatment mothers have a significantly higher probability of being able to enroll their children in childcare compared to control mothers.

In order for the estimated treatment effect to be unbiased, we need sorting by birthdate (into groups) to be random, so that the groups differ only in terms of kindergarten eligibility status. The selection of mothers into birthdate groups can be regarded random if the window around the cutoff is narrow enough: mothers of children born on December 31 can be assumed to be very similar to mothers of children born on January 1. However, the wider windows of 5 months around the cutoff used in our analysis - which are needed to ensure a large enough number of observations - mean that we need to consider certain possible sources of bias more carefully. Other age-related changes can lead to significant differences between the groups, because the average age of children in the two groups differs significantly. Figure 1 showed that in several countries, other significant changes occur around the child age at our

point of estimation: parental leave ends at age 3 of children in the CEE countries, and correspondingly, views regarding institutional childcare change as well. This means that due to the 5-month group windows, the treatment and control groups differ significantly in mean age when observed at a given point in time, and may be affected differently by these further factors in addition to the difference in their kindergarten eligibility.

To separate these other effects from the childcare effect, we define the estimation sample so that we include mothers in the treatment and control groups when their children are the same age. This sampling design ensures that child age, and therefore any further age-related characteristics - for example, child development or preferences regarding separation from the child - will be the same on average in the two groups. Table 4 summarizes the birth months included in the treatment and control groups, the months when each group is observed in the sample, and their age when they are observed for each cutoff used in the analysis.

*Table 4: Description of the birth and observation dates and child ages of the sample*

Cutoff	Birth months (treatment)	Birth months (control)	Enrollment date (treatment and control)	Observation months (treatment)	Observation months (control)	Child age at observation (treatment and control)
<b>Sept 1<sup>st</sup></b>	4-8	9-1	Sept 1 <sup>st</sup> (at age 3)	10-12	3-5	3y2m-3y8m
<b>Jan 1<sup>st</sup></b>	8-12	1-5	Sept 1 <sup>st</sup> (prior to age 3)	10-12	3-5	2y10m-3y4m
<b>March 1<sup>st</sup></b>	10-2	3-7	Sept 1 <sup>st</sup> (prior to age 3)	10-12	3-5	2y8m-3y2m
<b>May 1<sup>st</sup></b>	12-4	5-9	Sept 1 <sup>st</sup> (prior to age 3)	10-12	3-5	2y6m-3y
<b>May 1<sup>st</sup></b>	12-4	5-9	Sept 1 <sup>st</sup> (post age 3)	10-12	3-5	3y6m-4y

To estimate the causal effect of childcare availability on maternal labor supply, we turn to IV estimation, where treatment ( $T$ ) is an instrument for childcare availability. We estimate reduced form regressions separately for each country of the following form:

$$LFP_{cyi} = \beta T_{cyi} + \alpha_y + X'_{cyi}\pi_1 + \xi_{cyi} \quad (2)$$

where subscripts indicate yearly ( $y$ ), by country ( $c$ ) and individual ( $i$ ) variation, and  $LFP_{cyi}$  is the labor force participation dummy for individual  $i$ . The equation adjusts for a set of individual controls ( $X_{cyi}$ ), and  $\alpha_y$  represents year fixed effects. Controls include education, age, age squared, marital status, and region.

The parameter  $\beta$  captures the effect of belonging to the treatment group on the LFP probability. It can be interpreted as representing how much more active mothers are if they

are eligible for kindergarten rather than nursery school, which has significantly lower coverage. Since these rates differ by country, we interpret the magnitude of the childcare effect estimates based on their mean differences. This allows for a rudimentary analysis of the magnitude of the effects using a Wald estimator of the following form:

$$\beta^W = \frac{E(LFP_{cyi}|T_{cyi}=1) - E(LFP_{cyi}|T_{cyi}=0)}{E(C_{cyi}|T_{cyi}=1) - E(C_{cyi}|T_{cyi}=0)} = \frac{E(LFP^1) - E(LFP^0)}{E(C^1) - E(C^0)} \quad (3)$$

Since we do not directly observe enrollment in the EU LFS data, we proxy the country-specific childcare availability measures of the treatment and the control groups with the childcare enrollment rates of 3 and 2-year-olds respectively (reflecting country averages of kindergarten and nursery school enrollment rates).

In the setup presented so far, the treatment and the control groups differ in terms of both their dates of birth and dates of observation, which may introduce seasonal bias of various forms. First, the quarter of birth may be associated with various individual characteristics (Bound and Jaeger 1996). They cite a study that finds, for instance, that parents with higher incomes tend to have spring babies (Kestenbaum 1987). Second, child development may differ by season of birth, which may influence the mother's willingness to separate from the child. For instance, one study shows that health status and birth weight depend on the season of birth even after controlling for the characteristics of the mother (Currie and Schwandt 2013). The third possible bias is related to the different dates of observation of the groups. The seasonal variation of labor demand affects the actual and expected probability of employment, and thereby, the labor supply of mothers.

To remove possible seasonal bias from the measured effect, we estimate a second set of equations in which expand the sample with reasonably close labor market substitutes: mothers of children aged 4-6 years (separated into treatment and control groups based on the same cutoff date), and run a difference in differences (DID) regression. In the comparison sample, the treatment group as well as the control group has access to kindergarten, thus their childcare availability is the same. Thus, the comparison groups should be affected by the seasonal effects, but not the treatment effect, allowing us to separate out the seasonal factors. We construct a variable indicating the original and the comparison sample:

$$m_{cyi} = \begin{cases} 1 & \text{if } 3 \leq a_{cyi} < 4 \\ 0 & \text{if } 4 \leq a_{cyi} < 6 \end{cases} \quad (4)$$

where  $a_{cyi}$  indicates the age of the youngest child.

We then run the following reduced form regression separately for each country:

$$LFP_{cyi} = \beta^s T_{cyi} * m_{cyi} + \alpha_y + X'_{cyi} \pi_1 + \pi_2 T_{cyi} + \pi_3 m_{cyi} + \xi_{cyi} \quad (5)$$

where the estimated effect of treatment, corrected for seasonality, is given by  $\beta^s$ , the coefficient of the interaction term. It is important to note that, in addition to removing seasonal effects that are common to the main sample and the comparison group, results from this specification may also differ because it imposes a restriction on the model that the coefficients of further characteristics (controls) are the same for mothers of children of different ages. It may well be that the coefficients are, in fact, different for the original and the comparison sample of mothers, so the seasonality-corrected estimates may differ from the baseline estimates due to either seasonality biases being removed or the restriction on the other coefficients in the equation.

In each country, we run the estimation with the baseline and the seasonally corrected specifications, with and without controls. We measure the effect one quarter after the treatment (Q1), that is, in the quarter immediately after the September enrollment, as well as two quarters after the treatment (Q2). The Q2 results may represent longer-term effects, however, they may also be indicative of the flexibility of the September 1st enrollment date. For some countries, experts noted that enrollment is allowed year-round, depending on availability. It is therefore not possible to tell whether any significant effects observed in Q2 are due to longer term effects on maternal labor supply of enrollment in September, or shorter term effects due to enrollment later in the year.

Finally, we perform further robustness checks to verify the results. We estimate placebo cutoff regressions for mothers of 3-year-olds, based on other months besides the actual cutoff. As there should not be any sharp differences in childcare eligibility elsewhere, group membership (T) defined based on other months should not have an impact on participation. We also estimate the equations with the actual cutoffs, but for samples of mothers of children of different ages: at age 2, 4, 6, and 6. Again – once seasonality biases are controlled for using a comparison group – there should not be any significant effects, as childcare availability should not differ significantly by birthdate group at other child ages besides 3. Additionally, we re-estimate the seasonality corrected equations with comparison groups composed of mothers with different child ages. This should not affect the results, if seasonality biases that we want to control for are common to the different child age groups. The results of these tests point to the robustness of the results presented next.



## 4. Results

### 4.1. Childcare effect estimates by country

The estimates of the childcare effect around age 3 by country are presented separately for each group of countries in Table 5. The top panel in each table gives the estimates for treatment mothers observed in the quarter after the treatment (Q1), right after the September 1st enrollment date. The lower panel gives the estimates for the second quarter after the treatment (Q2), 3 months later. The relevant control groups are observed at dates that ensure that on average, the age of the youngest child be the same as it is in the treatment group. The first two columns within each country's results represent the baseline estimates, without and with controls, respectively. Here, the coefficient in our focus is that of the variable  $T$ , indicating the effect of having a child with a birthdate before the cutoff, and therefore being eligible for kindergarten (treatment). The next two columns present the seasonality corrected results, also without and with controls. Here, the coefficient of the interaction term  $mT$  represents the effect of kindergarten eligibility around age 3, net of any seasonal effects. The coefficient of  $m$ , the dummy indicating mothers with children around age 3, represents the difference in participation that is due to having a 3 years old child compared to having an older youngest child (aged 4-6). This is closely related to the idea of the relative labor supply rate discussed earlier: a high coefficient of  $m$  suggests that around age 3, a lot of mothers who are potentially ready to work have not yet returned to the labor market. The coefficient of  $T$  captures seasonal effects (birthdate effects) that are common to mothers of 3-year-olds and the comparison group of mothers with older children, which we want to exclude from the childcare effect estimates.

In the CEE countries (Table 5.a), the results generally point to a significant childcare effect of 0.07-0.11 in both Q1 and Q2. In the Czech Republic, we see a significant positive effect of around 0.1 in the baseline specification in Q1, which drops slightly to around 0.07 in the seasonality corrected estimates. There is also a positive childcare effect in Q2 of around 0.08 in the seasonality corrected case. Evaluating the results for Hungary in a similar manner, we can say that there a significant effect of around 0.08 in Q1, and an effect of around 0.1 in Q2. For Slovakia, we see strong evidence of an effect of around 0.09 in Q1 and 0.11 in Q2. The inclusion of controls does not significantly change the results in any specification, which supports the validity of  $T$  as an instrument. The coefficient of  $T$  indicates some significant age-independent seasonality for Hungary, but not for the other two CEE countries, and generally the seasonality correction affects the magnitude of the estimates slightly. Taken

together, the CEE results clearly point to a significant positive childcare effect of around 0.07-0.1 due to kindergarten eligibility around age 3 of children.

The results for the two Western EU countries (Table 5.b) in our sample provide some evidence of a positive childcare effect, as well as some differences between them. For Austria, we find no evidence of an effect in Q1, immediately after the treatment. For Q2, we find a strongly significant positive effect of around 0.07-0.08, which remains significant following the seasonality correction. For France, we find weaker evidence of an effect of around 0.05 in Q1, which loses its significance in the seasonality corrected specifications, and a similar magnitude estimate for Q2 that is just below significance. Neither Western EU country shows evidence of seasonality playing a significant role, the coefficient of T is very small, and insignificant. The estimates also do not change due to the inclusion of the controls. Overall, the results for Austria point to a significant positive effect of 0.07-0.08, and some suggestion of a smaller positive effect, but without strong evidence in the case of France. Any effect in western EU countries also does not appear to last as long as in the CEE countries.

The results from the Southern EU countries (Table 5.c) provide very little evidence of any childcare effect, and none once seasonality is controlled for. They show some significant estimates (0.05 for Greece and 0.08 for Italy) in Q2 in the baseline specifications, but insignificant estimates in all seasonality corrected specifications. The results imply that seasonality plays an important role in the maternal participation of these groups of mothers: the coefficients of T are highly significant and positive. This suggests that there are participation differences by birthdate groups that are common to mothers of 3 and 4-6-year-olds, and are therefore not due to differences in kindergarten eligibility. Once this effect is controlled for, the estimates suggest that for Southern EU countries, childcare availability around age 3 cannot be shown to have a significant impact on maternal labor supply.

Table 5.a: CEE countries

	Czech Republic				Hungary				Slovakia			
	Baseline		Seasonality Corrected		Baseline		Seasonality Corrected		Baseline		Seasonality Corrected	
Q1: 1 quarter after treatment												
T	0.0971*** (0.0320)	0.107*** (0.0318)	0.0246 (0.0252)	0.0346 (0.0261)	0.0328 (0.0297)	0.0283 (0.0288)	-0.0372 (0.0249)	-0.0406* (0.0229)	0.0614 (0.0376)	0.0636* (0.0384)	-0.0326 (0.0261)	-0.0296 (0.0261)
m			-0.598*** (0.0266)	-0.596*** (0.0273)			-0.301*** (0.0271)	-0.317*** (0.0265)			-0.545*** (0.0284)	-0.547*** (0.0286)
m*T			0.0724* (0.0407)	0.0678* (0.0408)			0.0700* (0.0387)	0.0770** (0.0369)			0.0940** (0.0458)	0.0965** (0.0456)
Constant	0.282*** (0.0204)	0.290 (0.520)	0.879*** (0.0171)	0.876*** (0.333)	0.438*** (0.0213)	0.0509 (0.371)	0.739*** (0.0168)	1.164*** (0.256)	0.323*** (0.0238)	0.0625 (0.510)	0.867*** (0.0156)	0.824** (0.329)
Controls		x		x		x		x		x		x
N	985	985	1,712	1,712	1,605	1,605	3,243	3,243	689	689	1,476	1,476
R <sup>2</sup>	0.011	0.071	0.316	0.340	0.001	0.099	0.074	0.184	0.004	0.054	0.271	0.295
Q2: 2 quarters after treatment												
T	0.0874** (0.0382)	0.0887** (0.0361)	0.0179 (0.0270)	0.0137 (0.0290)	0.0531* (0.0306)	0.0613** (0.0296)	-0.0258 (0.0283)	-0.0335 (0.0263)	0.0790** (0.0399)	0.0986** (0.0398)	-0.0121 (0.0339)	-0.00548 (0.0346)
M			-0.532*** (0.0297)	-0.527*** (0.0307)			-0.169*** (0.0295)	-0.193*** (0.0287)			-0.303*** (0.0351)	-0.306*** (0.0355)
m*T			0.0695 (0.0468)	0.0765* (0.0462)			0.0789* (0.0417)	0.0959** (0.0395)			0.0911* (0.0524)	0.107** (0.0524)
Constant	0.372*** (0.0243)	0.761 (0.554)	0.904*** (0.0170)	0.631* (0.373)	0.561*** (0.0223)	1.092*** (0.386)	0.730*** (0.0194)	1.398*** (0.273)	0.555*** (0.0277)	0.397 (0.579)	0.857*** (0.0216)	0.451 (0.381)
Controls		x		x		x		x		x		x
N	776	776	1,382	1,382	1,465	1,465	2,752	2,752	627	627	1,088	1,088
R <sup>2</sup>	0.008	0.121	0.266	0.309	0.003	0.107	0.020	0.132	0.006	0.065	0.084	0.120

Notes: Heteroskedasticity-robust standard errors in parentheses, stars indicate significance as:  $p < 0.01$ ,  $p < 0.05$ ,  $p < 0.1$ .

Table 5.b: Western EU countries

	Austria				France			
	Baseline		Seasonality Corrected		Baseline		Seasonality Corrected	
<b>Q1: 1 quarter after treatment</b>								
<b>T</b>	0.0360 (0.0303)	0.0483 (0.0310)	-0.00334 (0.0228)	-0.00834 (0.0227)	0.0503* (0.0288)	0.0491* (0.0281)	-0.00180 (0.0214)	-0.00422 (0.0208)
<b>m</b>			-0.114*** (0.0274)	-0.134*** (0.0275)			-0.107*** (0.0271)	-0.123*** (0.0268)
<b>m*T</b>			0.0394 (0.0380)	0.0570 (0.0372)			0.0521 (0.0359)	0.0535 (0.0355)
<b>Constant</b>	0.732*** (0.0220)	0.709 (0.482)	0.846*** (0.0163)	0.970*** (0.300)	0.779*** (0.0227)	-0.178 (0.470)	0.886*** (0.0148)	0.873*** (0.278)
<b>Controls</b>		x		x		x		x
<b>N</b>	1,043	1,043	2,220	2,220	1,140	1,140	2,389	2,389
<b>R<sup>2</sup></b>	0.002	0.048	0.015	0.064	0.004	0.082	0.015	0.082
<b>Q2: 2 quarters after treatment</b>								
<b>T</b>	0.0634** (0.0298)	0.0663** (0.0295)	-0.00765 (0.0273)	-0.0161 (0.0269)	0.0286 (0.0271)	0.0293 (0.0264)	-0.0283 (0.0274)	-0.0307 (0.0272)
<b>m</b>			-0.0872*** (0.0296)	-0.0956*** (0.0297)			-0.0584** (0.0274)	-0.0687** (0.0278)
<b>m*T</b>			0.0710* (0.0404)	0.0812** (0.0398)			0.0570 (0.0385)	0.0575 (0.0367)
<b>Constant</b>	0.747*** (0.0226)	0.659 (0.419)	0.834*** (0.0192)	1.066*** (0.285)	0.834*** (0.0203)	0.162 (0.513)	0.893*** (0.0184)	0.0375 (0.343)
<b>Controls</b>		x		x		x		x
<b>N</b>	996	996	1,948	1,948	898	898	1,630	1,630
<b>R<sup>2</sup></b>	0.006	0.044	0.008	0.044	0.002	0.155	0.004	0.126

Notes: Heteroskedasticity-robust standard errors in parentheses, stars indicate significance as:  $p < 0.01$ ,  $p < 0.05$ ,  $p < 0.1$ .

Table 5.c: Southern EU countries

	Greece				Italy			
	Baseline	Seasonality Corrected			Baseline	Seasonality Corrected		
Q1: 1 quarter after treatment								
T	0.0304 (0.0245)	0.0316 (0.0236)	0.0304 (0.0235)	0.0356 (0.0222)	-0.0207 (0.0773)	0.0335 (0.0695)	0.0905* (0.0537)	0.0887* (0.0511)
m			0.00263 (0.0234)	-0.00543 (0.0223)			0.00400 (0.0677)	-0.0332 (0.0636)
m*T			0.00001 (0.0340)	-0.00377 (0.0322)			-0.111 (0.0941)	-0.0940 (0.0851)
Constant	0.682*** (0.0174)	0.276 (0.382)	0.679*** (0.0156)	-0.233 (0.278)	0.639*** (0.0539)	1.679 (1.045)	0.635*** (0.0410)	0.965 (0.739)
Controls		x		x		x		x
N	1,662	1,662	3,477	3,477	235	235	671	671
R <sup>2</sup>	0.001	0.123	0.001	0.115	0.000	0.416	0.009	0.246
Q2: 2 quarters after treatment								
T	0.0486** (0.0243)	0.0507** (0.0237)	0.0562** (0.0262)	0.0604** (0.0253)	0.0847** (0.0386)	0.0798** (0.0353)	0.123*** (0.0363)	0.0801** (0.0325)
M			0.0106 (0.0247)	0.00137 (0.0234)			0.0191 (0.0353)	1.87e-05 (0.0322)
m*T			-0.00764 (0.0357)	-0.00893 (0.0341)			-0.0382 (0.0530)	-0.00697 (0.0476)
Constant	0.673*** (0.0170)	-0.261 (0.427)	0.663*** (0.0178)	0.344 (0.302)	0.561*** (0.0250)	0.154 (0.508)	0.542*** (0.0249)	0.387 (0.380)
Controls		x		x		x		x
N	1,739	1,739	3,202	3,202	1,048	1,048	2,177	2,177
R <sup>2</sup>	0.003	0.114	0.003	0.112	0.007	0.247	0.011	0.243

Notes: Heteroskedasticity-robust standard errors in parentheses, stars indicate significance as:  $p < 0.01$ ,  $p < 0.05$ ,  $p < 0.1$ .

To evaluate the cross-country differences in the childcare effect, we turn to a summary of some trends over the longer run. Figure 2 depicts the preferred seasonality corrected estimates with controls (red line), i.e. the coefficients of  $m * T$ , for the quarter before treatment (Q0), and four quarters after (Q1- Q4). Confidence intervals are included to assess their significance. The figure also depicts the relative labor supply of mothers of 3 year olds compared to those of older children (blue line), i.e. the coefficients of  $m$ , over time. No significant difference by birthdate is found prior to treatment, in Q0, in any of the countries, which supports the robustness of the cutoff-based method.

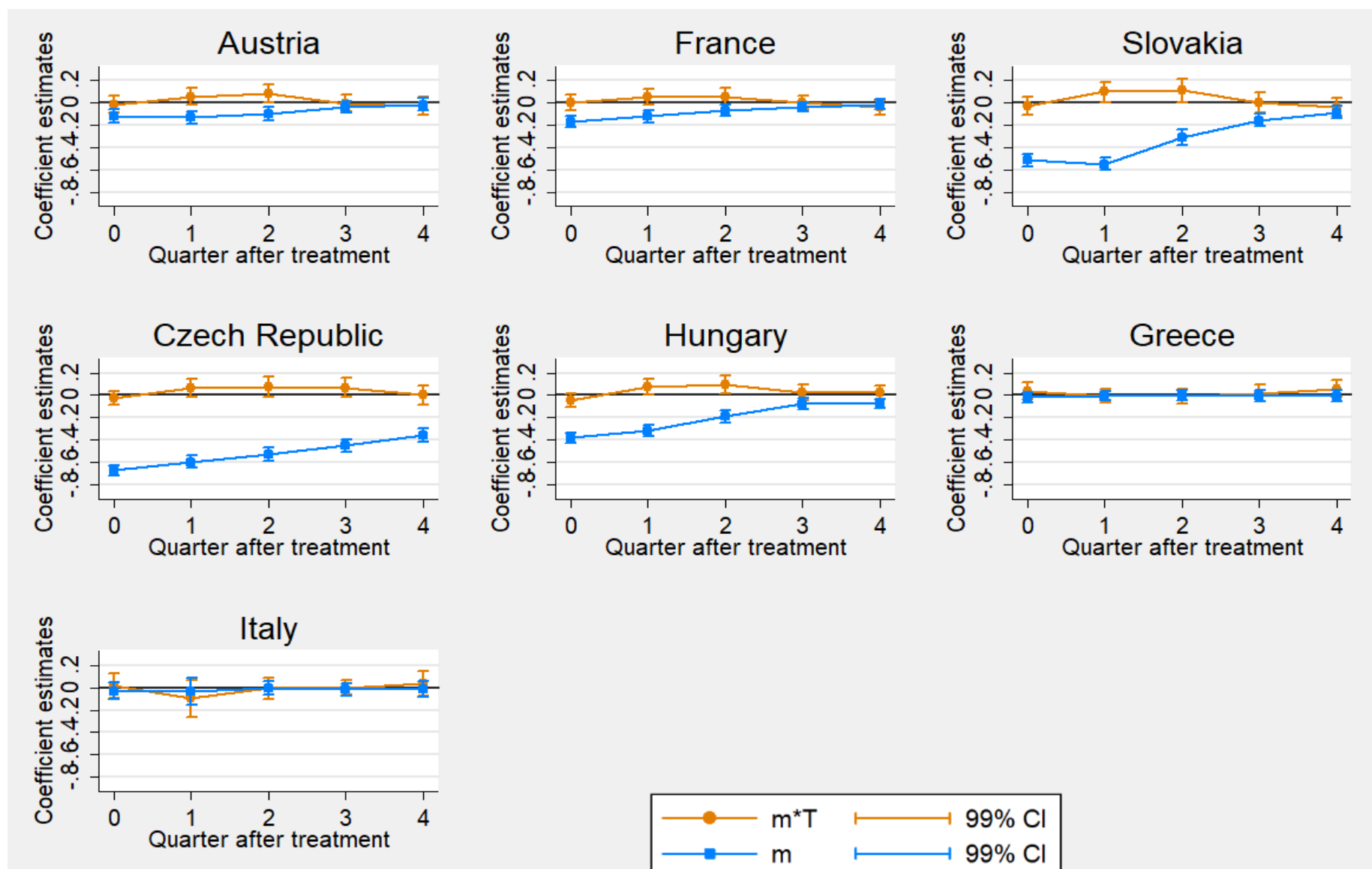
The long-run trend of the childcare effect estimates reflects what was discussed above: the highest, longest-lasting impact is seen in the case of the CEE countries. A somewhat

smaller and less significant impact is found for Western EU countries, though Austria more closely resembles the pattern of the CEE countries, while France shows only insignificant positive effects. No significant effect is seen in any quarter for the Southern EU countries. Looking at the relative labor supply of mothers of 3 year olds we see patterns in line with what we saw in Figure 1. The coefficients are significant and highly negative in the CEE countries, reflecting the high growth of maternal participation rates following age 3, and therefore, a high pool of mothers who are qualified and potentially ready to work at the point of estimation. The coefficients are still significant but much lower in magnitude in the Western EU countries, in line with some further growth in maternal participation after age 3. In Southern EU countries, the insignificant coefficients reflect the flatness of maternal participation over child ages seen earlier, suggesting that mothers who return to work after having a child have already done so prior to age 3 of their child.

The patterns seen in the coefficients of  $m$  are closely related to the findings regarding the childcare effect: it is the highest in countries where the relative labor supply of mothers at the given child age is still low, and lowest (or zero) in countries where it has already reached its long-run rate. In CEE countries, the labor supply of mothers of 3 year olds (our treatment and control group sample) is 0.4-0.6 below that of mothers of 4-6 year olds. Therefore, at the point of estimation, there is a large pool of mothers potentially ready and able to work when the treatment (kindergarten eligibility) takes place. In Western EU countries, the relative labor supply of mothers of 3 year olds is closer to, but still below that of mothers of older children (by around 0.2), reflecting the presence of a somewhat lower, but existing ready-to-work workforce. In the Southern EU countries, we do not see any significant difference in relative labor supply by child age, suggesting that by age 3 of children, mothers who return to work after having a child have already done so, and there is no significant ready-to-work maternal workforce present at the point of treatment.

The availability of a workforce that can be accessed in the short-run seems to be directly related to the childcare effect estimates found. Of course, the relative participation rate at a given child age is shaped by the institutional context relevant at and prior to the point of estimation. Furthermore, it is also important to consider the relevant incentives and constraints shaping the decisions of mothers. Therefore, we next discuss the role of the context in this light.

Figure 2: Dynamics of the effects



## **4.2. Cross country analysis: the role of institutions**

The interpretation of the magnitude of the childcare effect estimates is dependent on the difference in nursery school and kindergarten availability (coverage), i.e. the treatment effect. We therefore turn to statistics that take the treatment magnitude into account, shown in Table 6. It provides the childcare effect estimates, coverage rates used to calculate a proxy of the size of the treatment magnitude, and a Wald statistic indicating the magnitude of the childcare effect in each country that takes the magnitude (coverage difference between the treatment and the control group) into account. The coverage rates used in the calculation are meant to proxy the childcare availability of the groups as precisely as possible given data constraints: we use the enrollment of 2-year-olds in nurseries for the control group, and the kindergarten enrollment of 3-year-olds for the treatment group. Based on these, the difference in childcare availability, or treatment magnitude is the highest (around 0.4) in the CEE countries and Italy, somewhat lower in Austria and Greece (around 0.28), and the lowest in France (0.2). The table further summarizes the key institutional characteristics for each country for easier comparison.

The findings for the CEE countries provide strong evidence of a relatively large positive childcare effect. This is in line with our expectations, based on their institutional characteristics and the point of estimation (Table 1 and Figure 1). At this child age, besides maternal labor supply that is still relatively low, parental leave has just ended, so the financial incentives for staying home have decreased. At the same time, mothers can return to jobs that were still protected, making the return to the labor market easier. Cultural norms are also supportive of mothers' return to the labor market from around age 3, in line with the signals given by the institutional system. So, the institutional context at the child age where we estimate the childcare effect enables a large significant impact.

For the Western EU countries, the Wald statistics also point to a large childcare effect once treatment magnitude is considered, especially for Austria. These findings suggest that although the two Western EU countries – especially France – already exhibit higher maternal participation, childcare availability is still a factor that affects the labor supply of some mothers. Even at this age, there are mothers who are potentially ready and able to work, but constrained by the lack of childcare opportunities. For the Southern EU countries, the fact that we find no evidence of a childcare effect around age 3 can be explained by maternal labor supply that has already reached its limit by this child age. This pattern, in turn, is related to the



short length of leaves: job protection and financial leave benefits ended long before age 3 of their child, and mothers who were willing and able to work have already returned to the labor market. Furthermore, relatively traditional norms do not particularly support maternal employment even at older child ages. Childcare availability therefore has no impact at our point of measurement at age 3.

Keeping not only the country-level institutional context, but the point of estimation (age 3) in mind, it is important to note that the policy implications for expansion under and over age 3 are not straightforward. However, some important points can be made that are relevant to the evaluation of, for example, EU-prescribed childcare targets for under and over age 3 of children. Overall, we can say that childcare expansion is potentially the most effective if a pool of ready-to-work mothers is still present, and the timing of the end of job protection and leave payments in terms of child age coincides with the increase in childcare availability.

In CEE countries, further expansion above age 3 is likely to have a positive effect on maternal participation, as mothers are incentivized to return to work and yet some childcare shortage remains, though it is relatively low. The availability of childcare under age 3 is much further from the targets, and therefore subject to policy debates. Our estimates suggest that expansion has the potential to have a large impact due to the availability of a qualified workforce suggested by the maternal LFP rates at older child ages. On the other hand, our estimates pertain to a child age where leaves are just ending, and cultural norms change regarding whether mothers should stay home to care for their child. The effectiveness under age 3 may therefore be constrained by the long leaves and unsupportive cultural norms, as mothers may not be as willing to return, or encouraged by their environment to do so. Childcare expansion under age 3 should be coupled with a reform of the leave system, aimed towards shorter, better paid leaves that encourage greater paternal involvement, and the shaping of cultural views to be more open to institutional childcare under age 3 of children. Additionally, a greater availability of flexible, part-time work may also aid mothers who may be willing to separate from their child and return to work more gradually to decide to participate in the labor market.

Based on the results, childcare expansion in the Western EU countries may also have a significant positive impact, despite already relatively high participation rates. It appears that even at age 3, some mothers are effectively constrained by the lack of childcare opportunities. Expansion under age 3 is likely to have a larger impact because maternal participation is still below the rate of mothers with older children, and, depending on the country, cultural norms

are less likely to constrain the effectiveness. On the other hand, countries such as Austria – with relatively traditional views – must also shape views to avoid this constraint.

In the Southern EU countries, the potential effect of childcare expansion is limited by the low growth rate of maternal participation over child age, which is why we find no significant impact around age 3. At age 3, childcare does not appear to be the factor that effectively constrains maternal participation. Longer leaves with longer job protection periods coupled with childcare expansion may give more mothers an opportunity to return to the labor market after having a child. At the same time, the willingness and ability of mothers to return to work - as well as family policies themselves - are also affected by cultural views that are unsupportive of maternal employment, so changing these must also be a key element of effective policies.

Table 6: Childcare effect magnitudes and institutional environment

Country	$\beta^s$	p-value	Childcare enrollment (%) at child age ..		Difference in childcare availability	$\beta^w$	Maternal employment rate (%) at child age ...			Informal childcare	Child-related leaves			Labor market flexibility: Share of female part-time in employment (%)	Preferences / norms: Child suffers (%)
			2	3			0-2	3-5	6-14		Leave for mothers (weeks)	Total leave - average replacement rate (%)	Cash benefits at age 3 (%)		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)	(14)	(15)
Austria	0.08	0.02	26.5	54.7	28.2	0.29***	67	74	82	18.7	60	85.3	12.6	46	54.8
France	0.05	0.14	58.0	86.2	28.2	0.25	61	74	79	7.9	42	44.7	12.1	.	41
Hungary	0.08	0.04	16.7	60.2	43.5	0.19**	12	63	75	18.9	160	44.5	23	9	54.7
Czech Republic	0.07	0.09	7.0	41.9	34.9	0.20*	20	70	87	24.3	110	51.1	16.4	10	.
Slovakia	0.11	0.04	6.7	46.4	39.7	0.28**	15	56	80	16.3	164	32	17.7	6	44.4
Greece	-0.02	0.61	28.7	49.1	20.4	-0.07	50	54	59	32.5	43	53.9	5.2	12.5	65.3
Italy	-0.01	0.79	38.4	81.5	43.1	-0.02	51	53	56	18	48	52.7	5.2	31.6	61.8

(1) Seasonally corrected effect estimates from Table 6. Q1 or Q2 result included, depending on the effect measured.

(2) P-values of the seasonally corrected effect estimates from Table 6.

(3)(4)  $C^0$  and  $C^1$  Own calculations using EU-SILC data for 2- and 3-year-olds in formal childcare and pre-school services for years 2005-2012 based on the methodology of OECD Family Database. The data generally include children in center-based services, organized day care and pre-school (both public and private) and those who are cared for by a professional childminder, and exclude informal services provided by relatives, friends or neighbors. Exact definitions may, however, differ slightly across countries.

(4) Own calculations using EU-SILC data for years 2005-2012 based on the methodology of OECD Family Database.

(5) Difference in childcare availability between the treatment and the control group. ( $C^1 - C^0$ )

(6) Wald estimate = (1)/(5)

(7) Employment rate of mothers with youngest child aged 0-2, %. Source: OECD Family database, <http://www.oecd.org/els/family/database.htm> LMF1.2.C. Maternal employment rates by age of youngest child (2013)

(8) Employment rate of mothers with youngest child aged 3-5, %. Source: OECD Family database, <http://www.oecd.org/els/family/database.htm> LMF1.2.C. Maternal employment rates by age of youngest child (2013)

(9) Employment rate of mothers with youngest child aged 6-14, %. Source: OECD Family database, <http://www.oecd.org/els/family/database.htm> LMF1.2.C. Maternal employment rates by age of youngest child (2013)

(10) Own calculations using EU-SILC data for years 2005-2012 based on the methodology of OECD Family Database.

- (11) *Full-rate equivalent total paid leave for mothers (weeks). OECD Family database, <http://www.oecd.org/els/family/database.htm> PF2.1.A. Summary of paid leave entitlements (2015)*
- (12) *Average replacement rate (%): proportion of previous earnings replaced by the benefit over the length of the paid leave entitlement for a person earning 100% of average national earnings. OECD Family database, <http://www.oecd.org/els/family/database.htm> PF2.1.A. Summary of paid leave entitlements (2015)*
- (13) *Cash benefits and tax breaks at the child age of 3, relative to the median working age income, %. Source: Source: OECD Family database, <http://www.oecd.org/els/family/database.htm>*
- (14) *Part-time employment as a % of all employment, 20-64 year-old females, 2013. Data source: Eurostat, <http://ec.europa.eu/eurostat/data/database>*
- (15) *European Values Study, <http://www.europeanvaluesstudy.eu/>; Pre-school child suffers with a working mother. 0: Strongly disagree; 100: Agree strongly. Sample: 20-50 year-old females, waves 1999-2001 and 2008-2010*

## 5. Conclusion

This study estimates the effect of childcare availability on maternal labor supply for 7 European countries with different institutional contexts, and utilizes this variation to learn about the interdependencies of childcare and other factors. We provide comparable, quasi-experimental estimates – based on eligibility cutoffs – using harmonized data and a unified methodology. The results suggest that the childcare effect is the highest in CEE countries, where at this child age, maternal participation is still relatively low compared to that of mothers with older children, and paid leaves with job protection are just ending. We find some evidence of a smaller impact in Western EU countries, where, despite already high maternal labor supply around this age, some mothers who do so later on are still not working. In Southern EU countries, we find no significant childcare effect around age 3, which is due to the fact that leaves end at a much earlier age, so mothers who are ready and able to work have already returned to the labor market, and there is no further growth in maternal participation at older child ages. The impact of childcare expansion is dependent on the pool of available ready-to-work mothers, and the incentives and constraints faced by mothers at each given child age. The highest effect is shown for countries where the relative labor supply of mothers is still low, and where financial incentives (leave payments) and other constraints (lack of flexible jobs, unsupportive norms) are no longer present or binding.

Specific policy implications are derived from the results, considering the EU Barcelona targets for childcare expansion under age 3. For CEE countries, childcare expansion under age 3 has a high potential positive impact on maternal labor supply, however, it should be coupled with a reform of the leave system, aimed towards shorter, better paid leaves that encourage greater paternal involvement and the shaping of cultural views. In Southern EU countries, expansion has a lower potential impact due to many mothers permanently leaving the labor market after having a child. Longer leaves with longer job protection periods, coupled with childcare expansion under age 3 may give more mothers an opportunity to return to the labor market after having a child. Western EU countries may also have a significant positive impact, despite already relatively high participation rates, the magnitude of which is dependent on the relative labor supply of mothers of younger children.

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